Which predictor is the best to predict inflation in Europe: the real money-gap or a nominal money based indicator?

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Which predictor is the best to predict inflation in Europe: the real money-gap or a nominal money based indicator?[◊]

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Abstract
In the literature, two important views concerning the conduct of monetary policy are construed. One view is that the central banks' monetary policy must be credible if the authorities want to curb inflation. A second view is that central banks set their monetary policy by using all the information relevant for inflation and output projections. In Europe, a controversy has emerged about the role of monetary aggregates as useful indicators of future inflation and output. On one hand, evidence in favour of the usefulness of *nominal* monetary aggregates as good predictors is provided by the literature. On the other hand, empirical evidence in favour of *real* money indicators is found. The purpose of this paper is to contribute to the ongoing debate on the role of money aggregates in the setting of monetary policy. The question we are interested in is whether the real money gap contains more information about future inflation in Europe, than an indicator based on the growth rate of nominal money. We use a panel data framework instead of the usual time series methods on aggregate Euro data.

Keywords : monetary policy, inflation, panel data

JEL classification : E42, E52, E58, C23, C53

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1.- Introduction

In the literature, two important views concerning the conduct of monetary policy are construed.

One view is that the central banks' monetary policy must be credible if the authorities want to curb inflation. A wisdom argument is that credibility is gained through direct inflation targeting (DIT). This approach is defended in theoretical and empirical works.[1] The monetary authorities set the interest rates so that inflation and output do not deviate too much from target values. Implicit to this approach is the assumption that monetary policy cannot rely on intermediate monetary aggregates because money demand functions are not stable. DIT is an appropriate form of monetary policy when disinflation is the primary goal of stabilisation policy, which is the case in Europe. However, a fit of DIT models – for instance through the testing of Taylor rules - leaves an important percentage of the variance of the interest rate changes unexplained. This could mean that, in practice, this option is retained by a few European central banks.

A second view is that central banks set their monetary policy by using all the information relevant for inflation and output projections. In Europe, a controversy has emerged about the role of monetary aggregates as useful indicators of future inflation and output. On the one hand, evidence in favour of the usefulness of *nominal* monetary aggregates as good predictors is provided by the literature[2]. On the other hand, empirical evidence in favour of *real* money indicators is found[3]. The purpose of this paper is to contribute to the ongoing debate on the role of money aggregates in the setting of monetary policy. The question we are interested in is whether the real money gap contains more information about future inflation in Europe, than an indicator based on the growth rate of nominal money. We use a panel data framework instead of the usual time series methods on aggregate Euro data. The reasons for using panel data are two-fold. Firstly, the so-called "Euro-area" data are constructed using several aggregation procedures. Depending upon the conventions used, the data may display different characteristics[4], thereby implying no robust conclusions. Another reason why we prefer panel data is that, over the period under examination (1990-2004), the European countries can be

[1] See, among others, Bernanke *et al.* (1998), Clarida *et al.* (1998, 1999).

[2] See, Estrella and Mishkin (1997), Svensson (1999), Stock and Watson (1999).

[3] See, Nelson (2000), Trecoci and Vega (2000), Gerlach and Svensson (2003), Neumann and Greiber (2004).

[4] Some evidence for the European countries is provided by Eun-Pyo and Beilby-Orrin (1999).

considered as having divergent economic structures[5]. Panel data models help to capture such heterogeneities across the countries.

To bring evidence regarding the respective predictive power of the real money gap and a nominal money growth indicator for future inflation, we follow two approaches. We first consider an inflation equation containing the real money gap as a monetary indicator, along with two other macroeconomic variables that potentially influence future inflation: the output-gap and a component reflecting the inflation expectations. To model the private sector inflation expectations, we assume that the agents align their inflation forecasts on the central bank's implicit inflation objective. Then, this model is compared to another one where, instead of the real money gap, we consider the growth of the nominal money as a monetary indicator.

The rest of the paper is organized as follows. Section 2 briefly delineates the empirical money demand equation that is estimated in order to obtain a series of the real money-gap. Section 3 presents our estimates of the inflation model. Section 4 compares the predictive power of the real money gap and a nominal money growth indicator using out-of-sample forecasts. Section 5 concludes the paper.

2.- An empirical model of the real money demand

2.1.- The model

Our long-run equation represents a standard model of the demand for real money:

$$m_{it}-p_{it}=\alpha_0^i+a_1\pi_{it}+a_2y_{it}+a_3R_{it}^L+a_4R_{it}^S+a_5RER_{it}+a_6R_{it}^F+\varepsilon_{it}, \qquad (1)$$

where m_{it} is the nominal money stock, p_{it} is the price level, $\pi_{it}=4(p_{it}-p_{it-1})$ is the annualised inflation rate in quarter t, y_{it} is the real GDP, R_{it}^L is the long-run interest rate, R_{it}^S is the short-run interest rate, RER_{it} is the real effective exchange rate, R_{it}^F is the foreign interest rate and ε_{it} is a disturbance term. The index *i* refers to a specific country, while the index *t* refers to a quarter. All the variables written in lowercase are in logarithm, while the remaining others are in level; ε_{it} is independently and identically distributed among countries and quarters.

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[5] Although differences in economic structures were attenuated in the 1990's by the peer pressure mechanisms (Mastricht treasty and the Stability Pact) and the adoption of the single currency in 1999.

Theoretically, we expect the coefficients to have the following signs. The influence of inflation must be signed negatively ($a_1 < 0$), since an increase in inflation means a higher return on holdings of real assets. This is likely to induce a substitution between money and goods. The interest rate variables capture the impact of financial asset substitution. We expect the long-run interest rate to carry a negative sign (a_3<0) and the short-run rate to be positive (a_4>0). A rise in the long-run interest rate will lead to a decrease of the demand for money (to take advantage of higher returns on bonds), while an increase in the short term interest rate will result in a higher demand for money. A rising foreign interest rate is likely to translate into a decrease in the money demand (a_6<0), caused by a propensity to substitute away from domestic money. We expect the coefficient of real GDP to be positively signed (a_2>0). Finally, an appreciation of the real effective exchange rate results in an increase in the demand for domestic currency, so that we expect the coefficient a_5 to be positive (because an appreciation of the domestic currency is reflected by an increase in *RER* – see the description of the data below).

2.2.- The data

We use quarterly panel data spanning from 1990:1 to 2004:1. The empirical analysis focuses on the following group of the European countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, the Netherlands, Norway, Portugal, Spain, Sweden and the United Kingdom. The data are from several sources: the International Monetary Fund IFS database, the national central banks' statistics, the OECD main indicators and Eurostat database. The series are seasonally adjusted.

The nominal money stock, *m*, is computed as the logarithm of M3. The price level, *p*, is measured as the logarithm of the consumer price index. The inflation rate, π, is computed as the annualised inflation (that is $\pi_{it} = 4(p_{it} - p_{it-1})$). The long-run interest rate, R^L, is chosen as one of the following two variables (depending upon data availability): the Government bond yield or the bill rate. For the short-run interest rate, we use the money market rate as a proxy, or the commercial banks deposit rate when the money market rate series is not available. The real GDP, *y*, is the logarithm of the GDP at 1995 prices[6]. As a proxy for the foreign interest rate, we choose the US bond yield. Finally, the real effective exchange rate is

[6] The nominal money stock and the real GDP are expressed in US dollars.

measured as the trade weighted average of the real exchange rate based on bilateral trade shares and export similarity. An increase in *RER* reflects an appreciation of the domestic currency.

All the models are estimated over the period from 1990:1 to 1998:4. We use the data spanning from 1999:1 to 2004:1 to compute out-of-sample forecasts.

2.3.- Empirical results

To avoid the problem of spurious regressions, we begin our empirical investigation with an analysis of the unit root properties of the individual series. To this effect, we apply two tests based on Choi (2004) and Phillips and Sul (2003). They have some advantages over the conventional Im, Pesaran and Shin (2003) test (IPS) since they allow for cross-section dependence. The results in Table 1 indicate that we often conclude in favour of the unit root hypothesis, except for the inflation rate that is *I(0)*.

When estimating the demand for real money, we consider several combinations of the explanatory variables and this yields three models (models 1 through 3, which will be apparent in tables 2 and 3 below). We make use of estimators that exploit the information available in the cross-sections in order to obtain more precise estimates of the average parameters in the model. To this effect, we consider two types of panel estimators. Table 2 is made up of the results of long-run equations using a Fully Modified OLS (FMOLS) estimator. This estimator, proposed by Pedroni (2001), allows for heterogeneous slopes across the countries in addition to correcting for endogenous bias and serial correlation. Tables (3a) to (3c) also report some results based on mean group estimators as proposed by Pesaran and Smith (1995) and Pesaran *et al.* (1999). This allows to make explicit the speed of adjustment to the long-run equilibrium.

From Table 2, our main findings are as follows. The estimates match prior expectations from theory. The specifications show a positive income elasticity close to unity, which is consonant with the quantitative theory hypothesis. An increase in real income thus results in a proportionate increase in the demand for real money. The demand for money is negatively related to the rate of inflation. The order of magnitude of the semi-elasticity is small (around -0.3). The explanation for this is that low inflation in the Euro area prevents strong substitution effects from monetary assets to real assets or foreign currencies. The demand for real money is positively affected by the short-term interest rate and negatively correlated with the long-term interest rate. However, the semi-elasticity for the former is not statistically significant. This is surprising, but it may illustrate the fact that when one uses a broad definition of money

(M3 in our case) interest rates with the longest maturity better capture the effect of financial asset substitution. We also observe that the coefficient of the US bond yield is not statistically significant. The explanation of this can be found in the fact that, over the period under examination, the US interest rates have driven the European rates in a context of a 30-year trend of increasing integration across markets[7]. Thus, in the regression, the two rates are likely to be collinear. Eventually, we find the expected positive sign for the real effective exchange rate and this suggests a negative impact on money demand of a depreciating currency. This plays in favour of the currency substitution hypothesis.

As shown by the panel cointegration tests, a cointegration relationship between real money and the explanatory variables is accounted for. Indeed the *Panel* υ statistic is well above 1.64 and the other six statistics lie under –1.64. However, equation (1) assumes an instantaneous adjustment of the real money balance to its desired level. Such an instantaneous equilibrium state between the real money supply and the real demand for money is unlikely given the existence of transaction costs. We thus need to make a distinction between the long- and short-run behaviour in the money market by specifying an error correction mechanism of actual real cash balances toward their desired (long-run) level. In this view, Tables (3a) through (3c) report some estimations based on the mean group (MGE) and pooled mean group (PMGE) estimators. The three tables correspond to the models 1 through 3 respectively.

The PMGE method restricts the long-run coefficients to be equal across the countries (they are pooled), while the short-run coefficients are estimated individually and then averaged. The estimates are based on the application of the maximum likelihood approach and a Newton-Raphson algorithm to the following specification:

$$\Delta(m_{it}-p_{it})=\phi_i\,(m_{it-1}-p_{it-1})+\beta_i^{'}X_{it}+\sum_{j=1}^{p-1}\lambda_{ij}\Delta(m_{it-j}-p_{it-j})+\sum_{j=0}^{q-1}\mu_{ij}\Delta X_{it-j}+\eta_i+\varepsilon_{it}\,, \tag{2}$$

for each country i, where ε_{it} is the disturbance term and X_{it} designates the vector of the explanatory variables. If the roots of this equation lie outside the unit circle, then the model is error-correcting ($\phi_i<0$) and there exists a long-run relationship between the demand for real money and its determinants defined by

$$m_{it}-p_{it}=-(\beta_i^{'}/\phi_i)X_{it}+v_{it} \tag{3}$$

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[7] For recent empirical studies concerning the synchronization of the US and the European interest rates, see Hammersland (2004) and Chinn and Frankel (2005).

The long-run coefficients are equal to $\alpha_i = -\left(\beta_i' / \phi_i\right)$; the long-run homogeneity hypothesis is thus characterized by $\alpha = \alpha_i$ for every country i. Additionally, we calculate the mean group estimator which is an average of both the short- and long-run coefficients. We test for long-run homogeneity using a joint Hausman test based on the null hypothesis of equivalence between the PMGE and MGE estimator[8]. If the long-run coefficients are homogenous, then the MGE estimates are consistent and efficient (h-test in the tables). In the tables, the values of the *h-test* above 5% indicate acceptance of the null of poolability for the long-run coefficients. We can therefore consider that there are no significant statistical differences between the two estimators. The only exception is the coefficient of the real GDP (in model 3) for which the null of equivalent coefficients across the countries is rejected.

In all the regressions, the error-correction coefficients are very close to unity, implying that the short- and long-run coefficients are nearly equal. This suggests that the pressure on money demand to return to its long-run equilibrium is rather strong and that the adjustment time may be instantaneous. The short-term interest rate, which has no impact when we consider average-based long-run coefficients (the FMOLS or MGE estimator), does affect the demand for real money in the pooled-based regressions (PMGE). The corresponding coefficient is statistically significant either at the 10% or at the 5% level of significance with a semi-elasticity coefficient varying between 0.001 and 0.003. The US bond yield remains non significant. As for the other variables, we find that the real GDP positively affects real money demand, while both the inflation and the long-run interest rate have a negative influence (just as with the FMOLS estimator). Significant influences of the latter are found when we consider the PMGE estimator.

3.- An inflation model

Having estimated our long-run real money demand, we next study whether the real money-gap contains information about future inflation by specifying an inflation model. We assume that, in addition to the real money gap, inflation is also influenced by the output-gap and inflation expectations:

$$\pi_{it} = \mu_1 \pi_{it-1} + \mu_2 \hat{\varepsilon}_{it-1} + \mu_3 \left(y_{it} - y_{it}^*\right) + \mu_4 \pi_{i,t,t-1}^e + \omega_{it}, \qquad (4)$$

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[8] See Pesaran et al. (1996).

where $\pi^{e}_{i,t,t-1}$ is the expectation in quarter t-1 of inflation in quarter t by the private sector in country i, $\left(y_{it}-y^{*}_{it}\right)$ is the output-gap, $\hat{\varepsilon}_{it-1}$is the estimated residual from equation (1) and ω_{it} is a disturbance term.

We need to make some hypotheses about the way inflation expectations are formed. We assume a backward-looking mechanism and also consider that they are influenced by the implicit inflation objective set by the monetary authority:

$$\pi^{e}_{i,t,t-1}=\bar{\pi}_{it-1}+c(\pi_{it-1}-\bar{\pi}_{it-1}),\quad 0\leq c\leq 1, \tag{5}$$

where $\bar{\pi}_{it}$ is the implicit monetary authorities' inflation target. As a proxy of this target, we consider the "stable" component of past inflation computed by an HP filter. In the sequel, we call this variable the core inflation, and the difference $\pi_{it}-\bar{\pi}_{it}$ the inflation gap in quarter t for country i.

We also consider that the output-gap is endogenously determined as follows:

$$y_{it}-y^{*}_{it}=\upsilon_{1}\left(y_{it-1}-y^{*}_{it-1}\right)+\upsilon_{2}\left(R^{c}_{it-1}-\pi_{it-1}\right)+u_{it}, \tag{6}$$

where $R^{c}$is the Central Bank's interest rate which is set through a Taylor rule:

$$R^{c}_{it}=\alpha_{1}\left(\pi_{it}-\bar{\pi}_{it}\right)+\alpha_{2}\left(y_{it}-y^{*}_{it}\right)+v_{it}. \tag{7}$$

After a re-parametrization, the combination of (4) through (7) yields the following inflation equation:

$$\begin{aligned}\pi_{it}=\beta_{1}\,\pi_{it-1}+\beta_{2}\left(y_{it-1}-y^{*}_{it-1}\right)+\beta_{3}\left(\pi_{it-1}-\bar{\pi}_{it-1}\right)\\+\beta_{4}\left(\hat{\varepsilon}_{it-1}\right)+\beta_{5}\bar{\pi}_{it-1}+\eta_{it}\end{aligned} \tag{8}$$

where $\eta_{it}$is a disturbance term. This equation combines two aspects of monetary policy rules for inflation targeting: a targeting rule and an instrument-based rule. Deviations of the actual real money stock from its long-run target ($\hat{\varepsilon}_{it}$) signals a threat to price stability. An observed value above the target value may imply higher inflation, while the opposite situation may yield a deceleration of inflation. The Central Bank uses the information content of temporary disequilibria and reacts by changing its interest rate, given its implicit inflation objective and the output-gap. Equation (8) is estimated as follows. To deal with the endogeneity problem, we use instruments and apply a GMM estimator (instead of a standard IV estimator) to gain efficiency by exploiting additional moments restrictions. The t-statistics are computed using heteroscedastic- and serial correlation-consistent standard errors. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. These dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor. Finally, to avoid colinearity problems, we consider lagged inflation and core inflation separately in the regressors (instead of the inflation-gap). The results are shown in Table 4 (in the Table, the real money-gap is the estimated residuals from equation (1)).

As is seen, the coefficient of the lagged real-money gap is positive and statistically significant. We can interpret this finding as an indication of the predictive power of the real money-gap for future inflation. We note that the coefficient of the output-gap is significant (at the 10% level of significance) and that the core inflation is also significant at the 5% level of significance.

An interesting question is whether the real money-gap is more informative about future inflation than an indicator based on nominal money-growth. To address this question, we re-estimate equation (8) by substituting a nominal money growth gap indicator for the real money gap. More specifically, we replace $\hat{\varepsilon}_{it-1}$ by $\Delta m_{it} - \Delta m^*_{it}$, where Δm_{it} is defined as the annualised nominal money growth (similarly to the definition of the annualised inflation rate): $\Delta m_{it} = 4(m_{it} - m_{it-1})$. The target value of the nominal money growth, Δm^*_{it}, is computed as the fitted values (static forecasts) of the following equation:

$$\Delta m_{it} = \delta_1 \pi_{it-1} + \delta_2 \left(y_{it-1} - y^*_{it-1}\right) + u_{it} \tag{9}$$

This equation can be interpreted as the reduced form of a model that combines three equations, namely an Okun law equation, a Phillips curve and an aggregate demand equation. Given inflation, a higher output-gap results from an expansionary monetary policy ($\delta_2 > 0$). Besides, to achieve a lower inflation, the rate of nominal money growth needs to be adjusted downward ($\delta_1 < 0$). We estimate this equation using estimators similar to those used to estimate the real money demand and find two estimated coefficients that are statistically significant with $\hat{\delta}_1 = -1.13$ and $\hat{\delta}_2 = 2.83$. Once this is done, we compute the static forecasts and estimate the following equation by GMM:

$$\begin{aligned} \pi_{it} = {} & \beta_1 \pi_{it-1} + \beta_2 \left(y_{it-1} - y^*_{it-1}\right) + \beta_3 \left(\pi_{it-1} - \overline{\pi}_{it-1}\right) \\ & + \beta_4 \left(\Delta m_{it} - \Delta m^*_{it}\right) + \beta_5 \overline{\pi}_{it-1} + \eta_{it}. \end{aligned} \tag{10}$$

The results are shown in Table 4. We see that the coefficient of the nominal money growth-gap is statistically significant. To check the robustness of our estimations, we test for the validity of the instruments used in the GMM regressions. For each exogenous variable in the inflation equation, we first test for the joint statistical significance of the instruments. As is seen in Table 4, except for the inflation target, the instruments can be considered as good predictors of the exogenous variables. Indeed, the p-values of the Fisher test are less than 5% and this induces the rejection of the null hypothesis of no correlation between the explanatory variables and the instruments. Further, we apply the Sargan test to see whether the instruments can be considered as exogenous. As the p-values of the tests are higher than 5%, we conclude that the instruments are independent of the error term in the inflation equations.

4.- Out-of-sample forecasts

We now examine which of the two variables (the nominal money-growth rate or the real money-gap) is the most informative about future inflation. In this view, we perform out-of-sample predictions. We face some difficulties in doing such an exercise for the whole sample since, for some countries, there are no data for the money aggregate M3 from the period of the introduction of the Euro. For this reason, out-of-sample forecasts are computed for a selection of countries: Austria, Denmark, France, Italy, the Netherlands, Norway, Sweden and the United Kingdom. We consider two types of horizons relevant for monetary policy: short-term and long-term horizons. As an illustration of short-run predictions, we compute one-quarter ahead forecasts. For the longer time horizons, we consider one-year and two-year ahead predictions[9]. The forecasts are computed over the period from 1999:1 through 2004:1. We assess the stability of the relationships prevailing until 1998:4 for the period after 1999:1. Accordingly, the forecasts are based on the estimates presented in the preceding sections and on the actual values of the explanatory variables. More specifically, the forecasts are obtained from the estimation of the following equations:

$$\begin{aligned}\pi_{it}=&\beta_1 \pi_{it-j}+\beta_2\left(y_{it-j}-y^*_{it-j}\right)+\beta_3\left(\pi_{it-j}-\bar{\pi}_{it-j}\right)\\&+\beta_4\left(\hat{\varepsilon}_{it-j}\right)+\beta_5\bar{\pi}_{it-j}+\eta_{it}\end{aligned} \tag{11}$$

and

$$\begin{aligned}\pi_{it}=&\beta_1 \pi_{it-j}+\beta_2\left(y_{it-j}-y^*_{it-j}\right)+\beta_3\left(\pi_{it-j}-\bar{\pi}_{it-j}\right)\\&+\beta_4\left(\Delta m_{it-j}-\Delta m^*_{it-j}\right)+\beta_5\bar{\pi}_{it-j}+\eta_{it}\end{aligned} \tag{12}$$

where $j=1,4,8$. The estimated equations (11) and (12) are presented in Table 5.

[9] One-year and two-year ahead predictions cannot be obtained from equations (8) and (10); consequently, these equations have to be re-specified and re-estimated as equations (11) and (12).

In Table 6, we report the short-term horizon forecasts (one-quarter ahead forecasts). We compute the root mean squared errors of the predictions and discuss the predictive accuracy of the forecasts by examining a battery of test statistics. The numbers in bold correspond to the p-values of the cases where the inflation model with the nominal money growth rate yields forecasts that are better than those obtained with the real money-gap. As is seen, the predictions arising from the inflation equation with the nominal money growth rate are better than those coming from the equation with the real money-gap in the countries that gave up their national currency on January 1999 by adopting the Euro (France, Italy and the Netherlands)[10]. We obtain a similar conclusion for Denmark. The fact that, in the former three countries, nominal money growth dominates the real money gap in predicting inflation provides empirical support for the policies adopted by the European Central Bank since the adoption of the Euro. Indeed, the Eurosystem's monetary policy consists in maintaining price stability and choosing a nominal money growth indicator to achieve this objective (see European Central Bank (1999)). The result concerning Denmark can be explained by the fact that the Danish central bank pursues a monetary policy that ensures a stable Krone vis-à-vis the Euro. In the other countries, the predictions arising from the inflation equation with the real money gap are quite similar to those obtained from the equation with the nominal money growth. Considering long-term forecasts[11] (Table 7), the conclusions seem to be the same: the nominal money growth is still determinant in three out of four countries that use the Euro (Austria, France and the Netherlands).

So, our conclusion is that the nominal money-growth indicator has a higher predictive power of future inflation in the countries that adopted the Euro and where the monetary policy strategy is based on the one announced by the European Central Bank. In the other countries, monetary aggregates other than the nominal, for instance the real money gap, are equally informative for future inflation.

5.- Conclusion

This paper is an empirical contribution to the ongoing debate on the relative performance of a nominal money growth based indicator or the real money-gap in predicting future inflation. Our results provide support to the idea that the monetary policy in the Euro area countries is based on the choice of a nominal money indicator to achieve price stability. Such

[10] A different conclusion is obtained for Austria.

[11] We only report the results concerning the two-year horizon forecast, since those concerning the one-year forecasts were very similar.

a conclusion does not hold for the countries that continue to use their own currency. In the latter, the real-money gap may be equally informative about future inflation.

There is at least one direction in which this paper could be extended. In the time series domain, the empirical literature has shown that the money growth is highly correlated with the inflation rate at low frequencies[12]. It would be interesting to decompose the real money gap and the nominal money growth indicator into their high and low frequencies in order to identify the variable that plays the dominant role to explain short-run inflation and the one which is determinant for long-run inflation pressures. One major difficulty is, however, to transpose to a panel context the decomposition methods that use frequency domain approaches or other methodologies based on trend/cycle decompositions.

[12] See, among many others, Haug and Dewald (2004), Bruggeman *et al.* (2005), Assenmacher-Wesche and Gerlach (2006).

Table 1 – Panel unit root tests

		Choi(1)				Phillips and Sul (Z-test)(2)		
		Level	1^{st} diff	concl		Level	1^{st} diff	concl
Real money	P_m	-0.293	8.56	I(1)	c	-1.73	-3.35	I(0)
	Z	0.728	-5.74	I(1)	c,t	-0.35	-2.37	I(1)
	L*	1.238	-6.26	I(1)				
Real GDP	P_m	-0.88	3.85	I(1)	c	-3.33	-2.23	I(0)
	Z	2.05	-3.48	I(1)	c,t	0.06	-0.38	I(2)
	L*	2.32	-3.29	I(1)				
Inflation	P_m	4.67	-	I(0)	c	$-\infty$ (3)	-	I(0)
	Z	-2.96	-	I(0)	c,t	$-\infty$ (3)	-	I(0)
	L*	-3.33	-	I(0)				
Long-term interest rate	P_m	-0.277	16.68	I(1)	c	-0.55	-5.03	I(1)
	Z	0.833	-9.49	I(1)	c,t	2.08	-3.30	I(1)
	L*	0.94	-11.21	I(1)				
Short-term interest rate	P_m	-0.633	8.29	I(1)	c	0.99	-2.59	I(1)
	Z	0.66	-6.24	I(1)	c,t	1.29	-1.45	I(2)
	L*	0.68	-6.41	I(1)				
US bond yield	P_m	-0.27	16.68	I(1)	c	0.253	-6.97	I(1)
	Z	0.833	-9.49	I(1)	c,t	2.01	-5.81	I(1)
	L*	0.943	-11.21	I(1)				
RER	P_m	-0.10	4.88	I(1)	c	-1.42	-4.11	I(1)
	Z	0.109	-4.59	I(1)	c,t	-2.00	-1.74	I(0)
	L*	0.05	-4.44	I(1)				

Notes:

(1) All the statistics are distributed as standard normal asymptotically. The null hypothesis of a unit root is rejected for large positive values of the P_m statistic, while it is rejected for large negative values of the other two statistics. Accordingly, at the 5% level, we conclude as follows:

$$\textit{no unit root if} \begin{cases} P_m > +1.64 \\ Z < -1.64 \\ L^* < -1.64 \end{cases}$$

(2) The statistic is distributed asymptotically as standard normal. The null hypothesis of a unit root is rejected for large negative values of the Z-statistic. We thus conclude that the series does not have a unit root if the statistic reported is less than –1.64 (at the 5% level). c and c,t indicate that a constant term and a constant term plus a trend components are included in the regression.

(3) $-\infty$ means that we obtain a very large negative value.

Table 2. Pedroni FMOLS estimator – t-ratios in parentheses [(1)]

Endogenous variable: real money demand

	Model 1	Model 2	Model 3
Real GDP	0.969*	0.969*	0.900*
	(81.78)	(81.17)	(58.05)
Inflation	-0.268*	-0.260*	-0.295*
	(-5.26)	(-5.04)	(-6.11)
Long-term interest rate	-0.008*	-0.0082*	-0.0074*
	(-8.17)	(-7.89)	(-7.24)
Short-term interest rate	0.00074	0.0007	0.00003
	(1.82)	(1.64)	(0.84)
US bond yield	-	-0.0007	-
		(0.09)	
Real effective exchange rate	-	-	0.234*
			(6.25)
	Panel cointegration tests[(2)]		
Panel v	2.54**	2.56**	1.85**
Panel ρ	-4.20**	-2.80**	-3.22**
Panel *PP*	-12.96**	-11.97**	-13.39**
Panel *ADF*	-8.27**	-7.53**	-6.72**
Group ρ	-3.35**	-1.95**	-2.21**
Group *PP*	-14.00**	-12.84**	-14.06**
Group *ADF*	-9.62**	-8.68**	-8.36**

Notes:

(1) The FMOLS estimator is constructed by making corrections to the OLS estimator for the endogeneity of the regressors and serial correlation of the residuals. The endogeneity correction is achieved non-parametrically. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. The dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor problem. An asterisk (*) implies significance at the 5% level.

(2) All the statistics are distributed as standard normal asymptotically. The panel v rejects the null of no cointegration for large positive values (here for values higher than 1.64 at the 5% level) whereas the other six reject it with large negative values (here for values less than –1.64 at the 5% level). ** indicates the rejection of the null of no cointegration at the 5% level.

Table 3a. – Pooled Mean Group and Mean Group estimators – Model 1

Endogenous variable: real money demand

	PMGE[2]	MGE		
Long-run coefficients				
	Coefficients (t-ratios)	Coefficients (t-ratios)	h-test[1]	P-value
Real GDP	1.015* (50.519)	0.948* (14.165)	1.08	0.30
Inflation	-0.124* (-2.22)	-0.185 (-1.56)	0.34	0.56
Long-term interest rate	-0.005* (-4.00)	-0.009* (-2.12)	0.87	0.35
Short-term interest rate	0.003* (4.221)	-0.001 (-0.447)	2.44	0.12
Short-run coefficients[3]				
	Coefficients (t-ratios)	Coefficients (t-ratios)		
Real GDP	1.01* (45.90)	0.989* (10.08)		
Inflation	-0.124* (41.33)	-		
Long-term interest rate	-0.005* (-46.31)	-0.009* (-2.12)		
Short-term interest rate	0.003* (46.31)	-		
Δreal GDP(-1)	-0.073** (-1.762)	-		
Δinflation	-0.082** (-1.744)	-		
Error-correction coefficients	-0.995* (-47.38)	-1.026* (-30.08)		

Notes: * statistically significant at the 5% level of significance and ** statistically significant at the 10% level of significance

(1) h-test: Hausman test of poolability. The test is constructed as a test of equivalence between the pooled mean group and the mean group estimates. Probability values are provided for this test: p-values larger than 0.05 indicate acceptance of the null of poolability.

(2) The mean group estimates have been used as initial estimates of the long-run parameters for the PMGE. To deal with the problem of spatial correlation, prior to the estimation, we first regress the individual series on yearly time dummies and work with the residuals of these regressions. The dummies are intended to capture shocks that are shared across the different members of the panel and thus to remove a potential common factor problem.

(3) We report the short-run coefficients that are statistically significant (either at the 10% or the 5% level). The Schwarz coefficient has been used to select the lag orders.

Table 3b. – Pooled Mean Group and Mean Group estimators – Model 2

Endogenous variable: real money demand

	PMGE[2]	MGE		
Long-run coefficients				
	Coefficients (t-ratios)	Coefficients (t-ratios)	h-test[1]	P-value
Real GDP	0.991* (51.00)	0.922* (19.73)	2.65	0.10
Inflation	-0.226* (-4.17)	-0.210 (-1.703)	0.02	0.89
Long-term interest rate	-0.004* (-3.77)	-0.008** (-1.90)	0.74	0.39
Short-term interest rate	0.002* (3.091)	-0.001 (-0.597)	3.51	0.06
US bond yield	-0.000 (0.253)	-0.000 (-0.018)	0.41	0.52
Short-run coefficients[3]				
	Coefficients (t-ratios)	Coefficients (t-ratios)		
Real GDP	0.969* (57.0)			
Inflation	-0.221* (-55.25)			
Long-term interest rate	-0.004* (-55.72)			
Short-term interest rate	0.002* (55.72)			
Δreal GDP(-1)	0.032** (1.783)			
intercept	-0.001* (-7.29)			
Error-correction coefficients	-0.977* (-54.27)	-0.98* (-24.11)		

Notes: see footnotes of Table 3a

Table 3c. – Pooled Mean Group and Mean Group estimators – Model 3

Endogenous variable: real money demand

	PMGE[2]	MGE		
Long-run coefficients				
	Coefficients (t-ratios)	Coefficients (t-ratios)	h-test[1]	P-value
Real GDP	0.949* (38.54)	0.80* (12.57)	6.45	0.01
Inflation	-0.232* (-4.37)	-0.359* (-3.109)	1.53	0.22
Long-term interest rate	-0.002 (-1.608)	-0.004 (-0.966)	0.30	0.58
Short-term interest rate	0.001** (1.78)	-0.003 (-1.25)	2.80	0.09
Real effective exchange rate	0.134* (2.66)	0.337* (2.697)	3.15	0.08
Short-run coefficients[3]				
	Coefficients (t-ratios)	Coefficients (t-ratios)		
Real GDP	0.915* (38.12)	0.817* (15.64)		
Inflation	-0.224* (-37.33)	-0.406* (-3.10)		
Long-term interest rate	-0.002* (-38.10)	-		
Short-term interest rate	0.001* (38.10)	-		
Real effective exchange rate	0.129* (43.00)	0.409* (2.36)		
Error-correction coefficients	-0.965* (38.6)	-1.061* (-20.71)		

Notes: see footnotes of Table 3a

Table 4. – Estimates of the inflation model – GMM estimator with robust errors

	Inflation model with the real money-gap(1)	Inflation model with the nominal money growth gap(2)
Inflation (-1)	-0.10	-0.05
	(-0.67)	(-0.12)
Output-gap (-1)	0.279**	0.39**
	(1.74)	(1.77)
Core inflation (-1)	1.02*	0.94*
	(6.55)	(2.42)
Real money-gap(-1)	0.057*	-
	(2.45)	
Nominal money growth gap	-	0.02**
		(1.81)

Tests of the validity of the instruments: - Fisher statistics and p-values (in parentheses)

Regressor	Inflation model with the real money-gap	Inflation model with the nominal money growth gap
Inflation	3.55	3.55
	(0.007)	(0.007)
Output-gap	1.74	2.37
	(0.024)	(0.0008)
Core inflation	0.63	0.69
	(0.891)	(0.83)
Real money-gap	5.35	-
	(0.000)	
Nominal money growth gap	-	6.56
		(0.000)

Sargan tests: over-identifying restrictions – Chi-squared statistics and p-values

	Inflation model with real money-gap	Inflation model with nominal money growth gap
	20.78	20.27
	(0.236)	(0.260)

Note: t-ratio are in parentheses
* statistically significant at the 5% significance level.
** statistically significant at the 10% significance level
(1) List of instruments: inflation (lags 2 to 5), short-run real interest rate (lags 5 to 8), core inflation (lags 5 to 8), output-gap(lags 3 to 6), real money- gap (lag 7), inflation-gap (lags 8 to 11) .
(2) List of instruments: inflation (lags 2 to 5), short-run real interest rate (lags 5 to 8), core inflation (lags 5 to 8), output-gap (lags 5 to 8), nominal money growth gap (lag 8), inflation-gap (lags 8 to 11)

Table 5. – Estimates of the inflation model – GMM estimator with robust errors

1-year and 2-year-ahead inflation against the current values of the explanatory variables

	Inflation model with the real money-gap		Inflation model with the nominal money growth gap[(2)]	
	1-year[(1)] (j=4)	2-years[(2)] (j=8)	1-year[(3)] (j=4)	2-years[(4)] (j=8)
Inflation (-j)	-0.056	0.24*	-0.18	0.03
	(-0.48)	(2.09)	(-1.52)	(0.60)
Core inflation (-j)	0.857*	0.21	0.95*	0.57*
	(2.69)	(1.50)	(7.17)	(10.79)
Output-gap (-j)	0.827*	0.309*	1.11*	0.17
	(6.81)	(2.74)	(3.88)	(0.59)
Real money-gap(-j)	0.351*	0.05**	-	-
	(3.90)	(1.88)		
Nominal money growth gap(-j)	-	-	0.003	0.04*
			(0.194)	(3.54)

Tests of the validity of the instruments: - Fisher statistics and p-values (in parentheses)

Regressor	Inflation model with the real money-gap		Inflation model with the nominal money growth gap	
Inflation	2.93	4.02	2.81	3.53
	(0.0024)	(0.0002)	(0.007)	(0.0007)
Output-gap	1.54	1.78	1.80	2.06
	(0.042)	(0.035)	(0.038)	(0.03)
Core inflation	1.17	1.36	1.19	1.63
	(0.258)	(0.10)	(0.275)	(0.03)
Real money-gap	6.12	3.16	-	-
	(0.00)	(0.00007)		
Nominal money growth gap	-	-	3.57	3.01
			(0.00002)	(0.00002)

Sargan tests: over-identifying restrictions – Chi-squared statistics and p-values

	Inflation model with the real money-gap		Inflation model with the nominal money growth gap	
	33.40	38.14	15.58	27.67
	(0.121)	(0.09)	(0.11)	(0.07)

Note: t-ratios are in parentheses * statistically significant at the 5% significance level. ** statistically significant at the 10% significance level. *(1)* List of instruments: inflation (lags 5 to 12), core inflation (lags 5 to 12), output-gap (lags 5 to 12), real money-gap (lags 5 to 12). *(2)* List of instruments: inflation (lags 9 to 16), core inflation (lags 9 to 16), output-gap (lags 9 to 16), real money-gap (lags 9 to 16). *(3)* List of instruments: inflation (lags 6 to 12), nominal money growth gap (lags 6 to 12). *(4)* List of instruments: inflation (lags 10 to 17), core inflation (lags 10 to 16), output-gap (lag 10), nominal money growth gap (lags 10 to 16).

Table 6. – One-quarter ahead forecasts: RMSE and predictive accuracy tests (p-values in brackets)

Inflation model with real money-gap versus inflation model with nominal money growth

	Austria	Denmark	France	Italy	Neth	Norway	Sweden	UK
RMSE1[(1)]	0.80×10^{-2}	0.66×10^{-2}	0.16×10^{-1}	0.45×10^{-2}	0.156×10^{-1}	0.29×10^{-1}	0.14×10^{-1}	0.74×10^{-2}
RMSE2[(1)]	0.64×10^{-2}	0.72×10^{-2}	0.98×10^{-2}	0.48×10^{-2}	0.10×10^{-1}	0.28×10^{-1}	0.13×10^{-1}	0.75×10^{-2}

	AS[(2)]	SI[(2)]	WI[(2)]	NB[(2)]	MGN[(2)]	MR[(2)]
Austria	1.196	0.00	0.892	0.642	-1.253	-1.143
	(0.231)	(1.00)	(0.372)	(0.846)	(0.223)	(0.253)
Denmark	-1.929	-1.876	**-2.220**	1.202	1.799	1.155
	(0.053)	(0.06)	**(0.026)**	(0.332)	(0.085)	(0.247)
France	**7.348**	**1.876**	**2.463**	0.369	**-4.111**	**-2.277**
	(0.00)	**(0.06)**	**(0.014)**	(0.989)	**(0.0005)**	**(0.022)**
Italy	**-3.373**	**-2.293**	**-2.281**	1.175	1.829	1.413
	(0.0007)	**(0.022)**	**(0.022)**	(0.350)	(0.081)	(0.157)
Netherlands	**8.484**	1.042	**2.129**	0.429	**-3.822**	**-2.201**
	(0.00)	(0.297)	**(0.03)**	(0.976)	**(0.0009)**	**(0.027)**
Norway	1.106	0.208	0.821	0.951	-1.919	-1.186
	(0.268)	(0.834)	(0.411)	(0.547)	(0.068)	(0.235)
Sweden	1.319	-1.459	-0.912	0.961	-1.043	-1.217
	(0.1869)	(0.144)	(0.361)	(0.537)	(0.308)	(0.223)
UK	-0.871	-0.208	-0.152	1.014	0.407	0.433
	(0.383)	(0.834)	(0.879)	(0.486)	(0.687)	(0.665)

Note: (1)RMSE1 and RMSE2 are the root mean squared error of the predictions based respectively on equation (8) or (11) and equation (10) or (12).
(2) The different columns are: AS: asymptotic test, SI: sign test, WI: Wilcoxon's test, NB: naive benchmark test, MGN: Morgan-Granger-Newbold's test, MR: Meese-Rogoff's test. The null hypothesis is the hypothesis of equal accuracy of the predictions. The loss function is quadratic. The test statistics follow asymptotically different distributions: N(0,1) for the asymptotic test, the sign test, the Wilcoxon's test, the Meese-Rogoff's test, $F(T_0,T_0)$ for the naïve benchmark tests and a t_{T_0-1} for the Morgan-Granger-Newbold's test (where T_0 is the number of predicted observations). The Meese-Rogoff test statistic is computed with the Diebold-Rudebusch covariance matrix estimator. The truncation lag is 10 for the asymptotic test and is given by the integer part of $T_0^{4/5}$ for the Meese-Rogoff's test. For a detailed presentation of the different tests, the reader can refer to Diebold and Mariano (1995).

Table 7. – Two-year-ahead forecasts: RMSE and predictive accuracy tests (p-values in brackets)

Inflation model with real money-gap versus inflation model with nominal money growth

	Austria	Denmark	France	Italy	Neth	Norway	Sweden	UK
RMSE1[1]	0.11×10^{-1}	0.82×10^{-2}	0.16×10^{-1}	0.55×10^{-2}	0.19×10^{-1}	0.33×10^{-1}	0.17×10^{-1}	0.80×10^{-2}
RMSE2[1]	0.53×10^{-2}	0.71×10^{-2}	0.10×10^{-1}	0.47×10^{-2}	0.12×10^{-1}	0.33×10^{-1}	0.17×10^{-2}	0.72×10^{-2}

	AS[2]	SI[2]	WI[2]	NB[2]	MGN[2]	MR[2]
Austria	**2.74**	**2.324**	**2.55**	0.211	**-3.55**	**-2.08**
	(0.006)	**(0.020)**	**(0.01)**	(0.997)	**(0.003)**	**(0.037)**
Denmark	1.658	0.00	0.258	0.752	-1.752	-1.613
	(0.097)	(1.00)	(0.796)	(0.712)	(0.100)	(0.106)
France	**3.809**	**2.00**	**2.12**	0.435	**-2.41**	-1.69
	(0.0001)	**(0.034)**	**(0.034)**	(0.946)	**(0.029)**	(0.09)
Italy	0.840	1.00	1.499	0.741	-1.392	-1.238
	(0.401)	(0.317)	(0.133)	(0.721)	(0.183)	(0.215)
Netherlands	**5.499**	**2.50**	**3.05**	0.381	**-5.167**	**-2.29**
	(0.000)	**(0.012)**	**(0.002)**	(0.968)	**(0.0001)**	**(0.02)**
Norway	0.849	-1.50	-0.258	0.982	-0.253	-0.315
	(0.395)	(0.133)	(0.796)	(0.513)	(0.803)	(0.752)
Sweden	1.653	-0.50	0.103	0.909	-1.152	-0.856
	(0.098)	(0.617)	(0.917)	(0.574)	(0.267)	(0.392)
UK	**2.377**	1.00	1.861	0.820	-1.908	-1.407
	(0.017)	(0.317)	(0.062)	(0.651)	(0.075)	(0.159)

Note: see footnote Table 6.

References

Assenmacher-Wesche, K. and S. Gerlach (2006), “Interpreting the Euro area inflation at high and low frequencies”, Bank of International Settlements, Working Paper n°195.

Bernanke, B.S., Laubach, T., Frederic, S.M. and S.P. Adam (1998), *Inflation targeting: lessons from the international experience*, Princeton University Press, Princeton.

Bruggeman, A., Gonzalo, C.M., Björn, F. and J. Sousa (2005), “Structural filters for monetary analysis: the inflationary movement of money in the Euro area”, European Central Bank, Working Paper n°470.

Chinn, M. and J. Frankel (2005), “The Euro area and world interest rates”, Working Paper, University of Wisconsin, USA.

Choi, I. (2004), “Combination unit root tests for cross-sectionally correlated panels”, forthcoming in *Econometric theory and practice: frontiers of analysis and applied research: Essays in honour of Peter C.B. Phillips*, Cambridge University Press.

Clarida, R. Gali, J. and M. Gertler (1998), “Monetary policy rules in practice: some international evidence”, *European Economic Review* 42, 1033-1067.

Clarida, R. Gali, J. and M. Gertler (1999), “The science of monetary policy”, *Journal of Economic Literature* 37, 1661-1707.

Davidson, R. and J.G. Mackinnon (1991), “Several tests for model specifications in the presence of alternative hypotheses”, *Econometrica* 49(3), 781-793.

Diebold, F.X. and R. Mariano (1995), “Comparing predictive accuracy”, *Journal of Business and Economic Statistics*, 253-263.

Estrella, A. and F.S. Mishkin (1997), “Is there a role for monetary aggregates in the conduct of monetary policy?”, *Journal of Monetary Economics* 40, 279-304.

Eun-Pyo, H. and H. Beilby-Orrin (1999), “Compilation Methodologies for Euro zone Aggregates”, *OECD Monthly Note of Economic Leading Indicators.*

European Central Bank (1999), “The stability-oriented monetary policy strategy of the Eurosystem”, *ECB Monthly Bulletin*, February, 29-46.

Gerlach, S. and L.E. Svensson (2003), “Money and inflation in the euro area: a case for monetary indicators?”, *Journal of Monetary Economics* 50, 1649-1672.

Hammersland, R. (2004), “Who was in the driving seat in Europe during the nineties, international financial markets or the BUBA?”, Working Paper n°2004/20, Norges Bank, Oslo, Norway.

Haug, , A. and W. Dewald (2004), “Longer-term effects of monetary growth on real and nominal variables, major industrial countries, 1880-2001”, European Central Bank, Working Paper Series n°382.

Nelson, E. (2000), “Direct effects of base money on aggregate demand: theory and evidence”, Working Paper, Bank of England.

Neumann, M.J.M. and C. Greiber (2004), “Inflation and Core Money Growth in the Euro Area”, Deutsche Bundesbank Discussion Paper 36/2004.

Pedroni P. (2001), "Purchasing power parity tests in cointegrated panels", *Review of Economics and Statistics*, 83, 727-731.

Pesaran, M.H. and R. Smith (1995), "Estimating long-run relationships from dynamic heterogeneous panels", *Journal of Econometrics* 68, 79-113.

Pesaran, M.H., Smith, R.P., and K.S. Im (1996), "Dynamic linear models for heterogeneous panels", in Matyas, L. and P. Sevestre (eds), *The econometrics of panel data: a handbook of theory with applications*, 2nd revised ed., Kluwer Academic Publishers, Dordrecht.

Pesaran, M.H., Shin, Y. and R. Smith (1999), "Pooled mean group estimation of dynamic heterogeneous panels", *Journal of the American Statistical Association* 94, 621-634.

Phillips, P.C.B, and D. Sul (2003), "Dynamic panel estimation and homogeneity testing under cross section dependence", *Econometrics Journal* 6, 1, 217-259.

Stock, J.H. and M.W. Watson (1999), "Forecasting inflation", *Journal of Monetary Economics* 44, 293-335.

Svensson, L.E. (1999), "Monetary policy issues for the Euro system", *Carnegie-Rochester Conference Series on Public Policy* 51, 79-136.

Trecoci, C. and J.L.Vega (2000), "The information content of M3 for future inflation", ECB Working Paper n°33.